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Outline

This paper analyses the effects of monetary policy decisions on inflation expectations of European consumers. Using a novel approach that does not assume unbiasedness of expectations, which makes use of survey data on expected future as well as perceived past price developments and allows for non-normal peakedness and asymmetry, we convert qualitative survey responses of consumers in various European countries into quantitative time series of inflation expectations. After checking the rationality of the constructed expectations measures, we investigate the effects of unanticipated movements in interest rates and inflation on inflation expectations across European countries. We inter alia seek to explore whether the reaction of consumers in countries with more credible central banks differs from the reaction of consumers in less credible countries.

Keywords: Inflation expectations, Survey data, Monetary policy

JEL codes: C31, C32, E58

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1 INTRODUCTION

Following the start of Stage 3 of Economic and Monetary Union (EMU), a crucial objective for the European System of Central Banks (ESCB) is the rapid acquisition and maintenance of credibility for achieving price stability. Indeed, the credibility of the central bank is probably the single most important factor determining whether the pursuit of an anti-inflation policy is associated with significant output and employment losses.[1] When the central bank lacks credibility, when the public does not believe that the central bank is going to do what it says it is going to do, expected inflation in the private sector will exceed the central bank's objective for inflation. These expectations will feed into the wage and price decisions of households and firms, causing some workers and businesses to price their goods and services too high. The attendant decline in employment and real activity complicates the environment for monetary policy, making the central bank's job more difficult. Thus the public's expectations of inflation need to be taken into account by the central bank when determining the stance of monetary policy, in order to ensure realisation of the final objective (Kydland and Prescott, 1977; Barro and Gordon, 1983). Moreover, central banks need to assess the credibility of their monetary policy on an ongoing basis, and a key to this assessment is knowing how the inflation expectations of the general public compare with the price stability objective pursued by the central bank. However, measures of expected inflation are also of interest by themselves, as forecasting inflation is a major task of any central bank. Measures of expected inflation will play an important role in any such exercise, given that what firms and households expect inflation to be over various horizons influences their wage and price decisions, thereby feeding into the measured inflation rate.

Broadly speaking, there are two approaches to gauging inflation expectations.[2] The first is to try to infer the expected infation rate from the prices of financial instruments (Bank of Canada, 1998; Mylonas and Schich, 1999). The primary advantage of looking at the prices of financial instruments is that these prices reflect the expectations of agents upon which they have

[1] The finding of Alesina and Summers (1993) that the high degree of independence of the Deutsche Bundesbank is associated with low inflation with no attendant cost in terms of greater output volatility is generally seen as reflecting the high degree of credibility of that institution.

[2] An alternative route is to construct an economic model that includes expectations as variables and certain assumptions about how these expectations are formed. Estimating and solving the model then generates projected values for, e.g., the inflation rate. In this case, empirical analysis of the expectations can be carried out only indirectly, and is conditional on the behavioural model. This means that conclusions concerning the expectations will not be invariant to the choice of the underlying behavioural model (Kuismanen and Spolander, 1994).

been willing to act. This forward-looking nature makes financial asset prices popular among central bankers (see Hördahl, 2000, and Angeloni and Rovelli, 1998, for recent examples). If, for example, both nominal and index linked bonds with identical risk, liquidity and maturity characteristics are traded, it is in principle possible to obtain a very accurate measure of expected inflation (Barr and Campbell, 1997; Schmidt, 1999). However, in reality index-linked bonds are uncommon, and where they are issued, they usually differ from nominal bonds in more than just the determination of their returns. Thus, it is usually necessary to make strong auxiliary assumptions to infer expected inflation from the prices of nominal assets, thereby clouding the information content of the expected inflation series thus derived.' The alternative approach is to simply ask a sample of the general public what they expect inflation to be over some specified time horizon by means of a survey of some sort.This direct approach has the advantage that a measure of expected inflation is obtained which is undistorted by any auxiliary assumptions. The primary drawback is that survey participants may not base their decisions on their survey responses. Moreover, the results of sample surveys are overly sensitive to sampling errors and to the specific questions posed (Chan-Lee, 1980).

This paper deals with the derivation and use of quantitative information on inflation expectations from qualitative survey data. We contribute in two ways to the existing literature. On a technical level, our measure of expected inflation is an extension of the method made popular by Carlson and Parkin (1977), requiring less restrictive assumptions.This enables us, inter *alia*, to test for rationality of inflation expectations in stead of imposing it ***a priori. In this*** respect we extend the analysis of Bakhsi and Yates (1998) to 13 European countries or regions.[4] Besides this rationality test, our empirical analysis explicitly investigates the effects of the standard assumption of normal price distributions, i.e. peakedness and asymmetry, thereby extending the analysis of Dasgupta and Lahiri (1992) to five-category qualitative responses. The second contribution lies in the realm of monetary policy. More specifically, following the suggestion of Goodhart (1997), we investigate the information content of the expected inflation measures in the context of movements in actual inflation and short-term interest rates. We seek to explore whether the inflation perceptions of European consumers react to these events, and

[3] Other problems include the unreliability of financial market indicators. Financial markets tend to over-react to shocks and are susceptible to herding and speculative phenomena, leading to time-varying risk premia that hinder the use of such measures for monetary policy purposes.

[4] Moreover, Bakhsi and Yates (1998) base their analysis on quantitative survey responses whereas we must rely on a quantification of qualitative survey responses.

whether the reaction of consumers in countries with more credible central banks differs from the reaction of consumers in countries with less credible central banks. We find that, across European countries, the expectation measures based on the normal distribution most frequently show long-term equilibrating behaviour with respect to actual future inflation. The normal model is also the one in which long-run unbiasedness of inflation forecasts of consumers is most common across countries. The usefulness for monetary policy of the survey-based measures of expected inflation is hampered by the fact that they do not seem to be causal determinants of inflation. Moreover, and counter to intuition, they do not seem to react in any systematic way to upturns in inflation and surprise movements in short-term interest rates.

The remainder of the paper is structured as follows. The next section describes the methodology used in deriving the measures of expected inflation. Section 3 then puts these measures to the test empirically, by investigating their rationality and consistency, and by conducting some monetary policy experiments. Section 4 concludes.

2 QUANTIFICATION OF SURVEY-BASED EXPECTED INFLATION

The selection process of the method we use for extracting measures of expected inflation is influenced by the particular form in which the survey data are presented. Following Simmons and Weiserbs (1992), Madsen (1996) and Papadia (1983), we make use of the survey conducted monthly under the aegis of the European Commision. In this survey, respondents are questioned only about the expected direction of change, and can choose between more than three (but a finite number of) categories when responding. More specifically, European consumers are asked the following questions regarding prices:

1. Is the price level now compared to 12 months ago a) much higher, b) moderately higher, c) a little higher, d) the same, e) lower?

2. Do you expect prices over the next 12 months a) to rise faster, b) to show a similar rise, c) to rise less fast, d) to stay the same, e) to decline?[5]

[5] Both questions also include a 'don't know' category. In what follows, we allocate the numbers of this category proportionally to the other response categories . See Visco (1984, pp. 30,32) for a discussion.

To quantify the qualititative survey responses, we use a variant of the method which was first employed by Theil(1952) and which was made popular in a seminal article by Carlson and Parkin (Carlson and Parkin, 1977). The modifications to the Carlson-Parkin or CP-method are described in more (technical) detail in Berk (1999). An informal description of our method runs as follows. Within a cross-sectional sample of size N surveyed at time t, each agent i is supposed to answer questions about the future behaviour of prices at time $t+12$ (in months) on the basis of a subjective conditional probability distribution. This distribution is conditional on the information set available to the consumer at t. Agents are then supposed to report that no change in the price level is expected if the expected future inflation rate falls within an interval centered around zero. Similarly, agents will report that no change in the rate of inflation is expected if their expectation falls within an interval centered on the price increase that they perceived to have occurred in the past 12 months. The boundaries of both intervals, denoted as the response thresholds, are to be determined by the data. The survey results can be regarded as N drawings from the aggregate population and we are able to derive expressions of this expected inflation rate, the standard deviation and the response thresholds as functions of the survey responses. These expressions are given in relation to the perceived inflation rate, i.e. the price rise that consumers perceive to have occurred in the past 12 months. In order to obtain actual values for these variables, it is necessary to make two additional assumptions. First, the form of the aggregate density function must be known completely (except for its first and second moments). Second, the perceived inflation rate over the last 12 months must be known. We will discuss each of these additional assumptions in turn.

The distribution of inflation expectations

In most empirical applications of the CP-method the normal distribution is used (Seitz, 1988).[7] Our method ensures that the aggregate distribution of price changes expected by consumers in the survey is in fact the sum of independent identically distributed random variables (generated by the subjective distributions, each having finite first and second moments). Then the Central Limit Theorem (Kendall and Stuart, 1977) applies, and the aggregate distribution is normal. The normality assumption is also convenient: its statistical theory has been

[6] See Berk (1999) for details.

[7] For details regarding this and the other distributions used in this paper, see Johnson and Kotz (1970).

extensively developed, it is completely specified by two parameters and it is extensively tabulated, which makes the necessary computations easier to handle. However, the normal distribution also has certain disadvantages. As noted by, inter *alia*, Batchelor (198 1, 1982), Visco (1984) and Foster and Gregory (1977), there are *a priori* reasons for assuming that the aggregate distribution of expectations is not symmetric.* More specifically, with an upward trend in prices, the distribution of answers might be skewed to the right (Theil, 1952). In fact, Carlson (1975), studying the distribution of individual inflation expectations in the US, finds two non-normal features: the survey responses are more centrally peaked and tend to be distributed asymmetrically (more specifically, there is positive (negative) skewness during high (low) inflationary periods).[9] Imposing a normal distribution then yields downward (upward) biased estimates of the expected inflation rate during high (low) inflationary periods (Smith and McAleer, 1990). The results of Pesaran (1987) and Pesaran and Wright (1991) seem to confirm this. Normality was also rejected by Batchelor (198 1) and Batchelor and Dua (1987) for French, German and UK data. However, Balcombe (1996), using consumer survey data from New Zealand, extends the CP approach to allow for skewness and kurtotic distributions but finds no empirical evidence in favour of such distributions. Conclusive empirical evidence concerning the form of the distribution is difficult to obtain, as it requires surveys for which the microresponses are not only available (which is often not the case for privacy reasons), but also quantitative in form. In light of the previous discussion, and without access to data on quantitative price changes of individual consumers, which would allow us to check the accuracy of alternative distributional assumptions, different types of distributions will be used and tested for their appropriateness. We begin by following existing practice and use the normal distribution. [10] In addition, we use a scaled-t distribution. There are several reasons for choosing this type of distribution. First, the *t*-distribution can be motivated in exactly the same way as the normal distribution. This is because our appeal to the Central Limit Theorem above was unnecessarily restrictive (Batchelor, 1981). In general this Theorem states that the sum of independent, identically distributed, random variables will tend to some particular class of distributions (the family of 'stable laws'), of which the normal distribution is but one example and of which the t-distribution is also a member.

[8] Also, the maximum negative outcome is by definition limited, whereas the maximum positive response is, in theory, unbounded. This also suggests skewness to the right.
[9] See in this respect also Roger (2000).
[10] Dasgupta and Lahiri (1992) also apply the logistic distribution, and find that logistic models are very similar in performance to the normal model.

Second, research by Carlson (1975) has shown that a symmetrically scaled t-distribution provides a better approximation to the distribution of inflationary expectations than the normal distribution, in particular because of its peakedness. [11] Third, it can be modified relatively easily to incorporate asymmetry, and a comparison of central-t and noncentral-t gives an indication of the effects of asymmetry on measures of inflation expectations. In sum, by using normal, central-t and noncentral-t distributions we allow for nonnormal peakedness as well as asymmetry to influence our expectation measures.

For the qualitative surveys which we are considering, however, there are not enough independent response categories to estimate other distributional parameters besides the mean and standard deviation. The degrees of freedom and the asymmetry-parameter therefore have to be determined outside the model. Regarding the degrees of freedom, we use (for the central-t as well as the noncentral-t distribution) the result of Carlson (1975), who finds that a number equal to six gives the best estimates.'* With respect to the asymmetry parameter relevant for the noncentral-t, we assume that the degree of asymmetry varies directly with the level of inflation: as the inflation rate increases (decreases), the proportion of consumers expecting an increase in future prices will increase (decrease) as well. We use two measures of asymmetry, both of them time-varying. In the first measure, the degree of asymmetry is measured as the difference between the last official inflation rate available to consumers when responding to the survey, and the average of (twelve month) inflation rates in the previous 12 months. [13] The second measure follows Batchelor (198 1), and uses the deviation of actual inflation during each month from its mean over the whole sample period.

The perceived inflation rate

As shown formally in Berk (1999), the perceived past inflation rate performs a scaling role with respect to the expected future inflation rate. A possible proxy for this perceived rate is the most

[11] The preference of Carlson (1975) for the scaled t-distribution was based on empirical, not theoretical, grounds. It turned out that this distribution provided the best fit to the observed subjective distributions of respondents to the Livingston survey in the US.

[12] When comparing different distributions, a low number of degrees of freedom seems preferable, since a scaled t-distribution can be made arbitrarily close to the normal by increasing the number of degrees of freedom.

[13] i.e. $asym(t) = \Pi(t-1) - \sum_{i=1}^{12} \Pi(t-i-1) / 12$, thereby assuming that consumers have a horizon of 12 months when using past information, and that the inflation outcome in each of these months has an equal influence on the consumers' expectations.

recent inflation rate available to consumers when answering the survey question regarding future prices (Simmons and Weiserbs, 1992). More specifically, given the publication lag involved this would imply: $\Pi_t^p = \Pi_{t-1} = \ln(cpi_{t-1}) - \ln(cpi_{t-13})$. However, there is no *a* priori reason to expect that the inflation rate perceived by consumers is adequately and completely represented by official inflation figures. In fact, due to the 'signal processing problem' described in Lucas (1972, 1976), respondents to the survey, even if behaving rationally, may not perceive correctly the aggregate inflation rate. An alternative measure of the perceived inflation rate makes use of survey information, i.e. the answers of consumers to the question pertaining to price developments in the past 12 months. But we have to take into account the fact that the two questions regarding prices in the survey are phrased differently. We therefore transformed the first question by aggregating the responses pertaining to different intensities of inflation (i.e. prices in the past 12 months are 'much higher', 'moderately higher' and 'weakly higher') into 'positive inflation'. Together with the categories 'prices remained the same in the past year' and 'prices went down in the past year' we are left with the 'traditional' survey categories for which the CP-method was originally developed. The responses thus allow us to calculate the mean value of the perceived past inflation rate and the standard deviation of this rate, contingent on the form of the distribution function and a scaling factor representing a response threshold. The latter was chosen so as to equate the mean of the perceived past inflation rate with the mean of the actual past inflation rate.[14] This assumption differs from the one made by CP, where the unbiasedness assumption applied to expectations of future price developments. The latter assumption is especially problematic if the expectations data are to be used in tests of theories of expectations formation, such as the rational expectations hypothesis. Another advantage of the proposed extension is that it allows for time-varying response thresholds, in contrast to the original CP-method. This relaxation of some important assumptions of the original CP-method makes our extension *a* priori interesting. However, before proceeding to the application of our method, some caveats should be mentioned. First, there is no reason why the response thresholds should be symmetric, as we assume. Second, it is assumed that the process of formation of the expected future inflation rate is independent of the formation of the perceived

[14] The implication is that the perceived inflation rate is an unbiased estimate of the actual inflation rate. However, this does not imply that the expected future inflation rate, which is the product of the perceived rate and a nonlinear term representing the frequencies of the assumed cumulative distribution function (see Berk, 1999), is by construction an unbiased estimate of the inflation rate.

past inflation rate. It is unclear how restrictive this assumption is, and investigating the consequences of possible interactions seems an interesting topic for future research.

3 Empirical analysis

In our empirical work, we use survey data obtained from the European Commission. The data are monthly and seasonally adjusted, covering the period January 1986 up to December 1999. [15] The data pertain to the countries comprising the EU (excluding Luxemburg) and two regions: the euro area (ie the 11 European countries which have adopted the euro as their currency as of January 1999) and the EU. Upon prior investigation of the data, we eliminated Austria, Sweden and Finland from our sample because of insufficient observations. The survey responses are complemented with data on consumer price inflation, calculated as the increase of the CPI over the previous corresponding month. As announced, we will distinguish between methods using a normal distribution, symmetric scaled t-distribution, a noncentral scaled t-distribution based on an asymmetry parameter using only information available to consumers when responding to the survey and a noncentral t-distribution based on an asymmetry parameter using information based on the entire sample. The symmetric distributions are calculated using both the most recent inflation figure available to consumers (ie the inflation rate of the previous month) and the inflation rate which consumers perceived to have prevailed in the past 12 months as scaling parameters. [16]

Consistency and rationality

We start by investigating the time series properties of the constructed expectations measures as well as the actual inflation rate. It is important to note that possible nonstationarity of the former is not at odds with the method used to construct the expectations measures, which assumed

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[15] Only seasonally adjusted data are available from the European Commission. The weighting of individual countries in the construction of the data for the euro area and the EU is done by the European Commission, and discussed in European Economy, Supplement B.
[16] The latter option was excluded for the asymmetric distributions, since it is not consistent with the assumptions on which the asymmetry parameter was based.

constant first and second moments. This is because our method is based on a cross-sectional sample of size N surveyed each month, so that the construction of the expected inflation measures is not influenced by possible persistence. The same, however, does not apply to the constructed *time series* of expected inflation rates.

[Insert tables 1 and *2* here]

Tables 1 and 2 present the results of the integration analysis following the procedure recommended by Dickey and Pantula (1987), who emphasize the possibility of multiple unit roots. They show that first testing for *k* unit roots, and upon rejection proceeding with testing for *k-l* unit roots, and repeating this procedure until a null is accepted, ensures consistency. The tables overwhelmingly indicate that the inflation rate has a single unit root (that is, the price level is I(2)). The same applies to the measures of expected inflation and inflation uncertainty (ie the standard deviation of the expected inflation rate as estimated from the survey data). [17] This has consequences for the admissable econometric strategy used in investigating the information content of expected inflation rates. Regressing actual inflation rates on expected inflation rates for example, as is done in the literature on the rationality of inflation expectations derived from survey data (see Pesaran, 1987, for an overview) would (in the absence of cointegration, see below) yield inconsistent parameter estimates, whether or not the covariance matrix is corrected parametrically to adjust for autocorrelation and heteroskedasticity. [18]

Given the high degree of persistence of the series under consideration, we next investigated the possibility that expected and actual future inflation rates (ie the inflation rate prevailing in the coming 12 months) are cointegrated. The concept of cointegration, which stresses long-term relationships, seems a suitable methodology given the orientation of monetary policymakers, who frequently stress the medium- to long-term horizons when striving for price

[17] The exception being the expected inflation rate of Denmark, for which the null of stationarity could not be rejected.

[18] Such a correction is necessary given the fact that the forecast horizon (in our study 12 months) exceeds the sampling interval (1 month in our case) (See Brown and Maital, 1981; Papadia, 1983). OLS in this case generates consistent parameter estimates, but inconsistent estimates of the covariance matrix, because, in violation of the classic OLS-assumptions, the disturbances are not serially uncorrelated but follow an MA(11) process. See Hansen and Hodrick (1980), Hansen (1982) and Hamilton (1994) for details.The Newey-West (1987) standard errors are asymptotically consistent in the presence of autocorrelation as well as heteroskedasticity.

stability (Bemanke and Mishkin, 1997). Cointegration implies that, although both the actual 12-month-ahead and expected inflation rates show substantial persistence and show no mean-reverting behaviour, both series form an equilibrium relationship in the sense that deviations from this relationship are temporary. This cointegrating relationship can, under certain additional conditions specified below, be of use for monetary policy purposes. Table 3 presents the results of the cointegration analysis, using the techniques developed by Johansen (see Johansen, 1995, for details).

[Insert table 3 here]

The analysis presented in table 3 is based on the Johansen-type likelihood ratio tests, with the dimension of the VAR on which these tests are based in parentheses. Based on a visual inspection of the data, we decided against assuming deterministic trends in the data when constructing the VAR. Regarding the specification of the lag order of the VAR, a general-to-specific methodology is followed. Starting from a maximum lag length of 12 months, we used information criteria (AIC and SBC) and likelihood ratio tests to determine the appropriate lag length. It becomes clear from table 3 that the expected inflation rates based on symmetrical distributions are more often cointegrated (ie for more countries) with the actual inflation rate 12 months ahead than the asymmetrical ones.[19] Moreover, from a cointegration point of view, the expectations measures scaled with the actual inflation rate fare better than the ones scaled with the inflation rate that consumers perceived to prevail in the past 12 months. Finally, there seems little to choose (in terms of evidence of cointegration) between the normal and scaled symmetric-t distributions. From a country perspective, only in France, Italy and Portugal no evidence of cointegration could be found.[20] The strongest evidence, defined in terms of the number of different expectations measures for which cointegration with the actual future inflation rate could be established, was located in Belgium, Germany, the Netherlands and the UK.

[19] These results, of course, do not allow the conclusion that asymmetric distributions, in general, perform worse than symmetric ones. Our findings could be due to, inter alia, the particular selection of the number of the degrees of freedom or of our choice of asymmetry parameter.
[20] Denmark was excluded from the analysis because of the stationarity of the expected inflation rate.

The cointegration framework can be used to investigate the forecast consistency of inflation forecasts, as defined by Cheung and Chinn (1997). A forecast is consistent if it has a one-to-one relationship with future inflation in the long run (see also Fischer, 1989). This concept involves the behaviour of the forecast relative to the actual, over time and on average. The forecast consistency property implies that expected and future inflation (i) have the same order of integration, (ii) are cointegrated and (iii) have a cointegration vector $\Pi^e(t) = a + b\Pi(t + 12)$ in which the restriction $a = 0$, $b = 1$ applies. It follows from tables 1 to 3 that the requirements (i) and (ii) are fulfilled for a large number of expectation measures. We investigated restriction (iii) by testing for the existence of a unit root in the forecast error $\Pi^e(t) - \Pi(t + 12)$, with the Augmented Dickey Fuller test implemented in a sequential manner as prescribed by Dickey and Pantula (1987).

The motivation for using this procedure instead of restricting the cointegrating vectors obtained by the Johansen procedure in table 3 is twofold. First, using the available prior information on the coefficients implies an efficiency gain in the procedures for investigating the time series properties of the data *vis-à-vis* not using this information. Second, the Johansen procedure assumes normality of the residuals. However, many of the residuals do not pass the (Jarque-Bera) normality test.[21] Huang and Yang (1996) conclude that the Engle-Granger method is more robust and dependable than the Johansen approach in the absence of normally distributed residuals. Investigating the stationarity of the forecast error implies implementing the Engle-Granger cointegration procedure, subject to the parameter restriction mentioned above. As can be seen from table 4, we could reject the null of 2 and a single unit root for most measures of expected inflation forming a cointegrating relationship with future inflation, thereby confirming the rejection of non-cointegration implied by the Johansen procedure. Moreover, restriction (iii), which implies long-run unbiasedness of the inflation forecasts of consumers, therefore seems not to be refuted by the data. This is in line with the findings of Batchelor (198 1) and Papadia (1983).

[insert table 4 here]

[21] Results are available from the authors on request.

It should be noted that the concept of forecast consistency focusses on the long-run property of forecasts, and hence is weaker than the one conventionally used in evaluating forecast rationality. It does not impose any further restrictions on the forecast errors, over-and-above the requirement that they be weakly covariance stationary.[22] As our survey forecasts of expected inflation are very likely to be subject to measurement errors (Smyth, 1992), this concept of forecast consistency is especially useful, since it allows for serially correlated forecast errors. The latter can happen, for example, when stationary measurement errors are present (see Lee, 1994; Cheung and Chinn, 1997, for details).[23] However, the possibility of nonstationary measurement errors, as well as the relatively small sample size on which the (co)integration tests are based, are important caveats to be kept in mind when interpreting the results.

In order to keep the subsequent empirical analysis on monetary policy implications tractable, we need to select a measure of expected inflation from the 6 available alternatives. Based on the results presented in tables 3 and 4, we conclude that the expectations measured based on the symmetrical distributions scaled with the actual inflation rate available to consumers when responding to the survey outperform the other alternatives.[24] In order to choose among the remaining alternatives, we checked the accuracy of the expectations measures. Table 5 documents the mean abolute prediction error (MAPE), the root mean square prediction error (RMSE) and Theil's inequality coefficient (Theil) of the forecast error derived from expectations measures based on the normal and symmetric t-distributions (both scaled with the lagged inflation rate), respectively. It follows from table 5 that the accuracy of the expectation measure based on the normal distribution slightly exceeds the one based on the symmetric-t distribution. We therefore proceed with the expected inflation rate based on the normal distribution as our preferred measure of expected inflation when using the generated inflation expectations measures for monetary policy purposes.

[insert table 5 here]

[22] In addition to unbiasedness, other necessary conditions for rationality include efficiency and orthogonality. Both conditions imply forecast errors that are, in general, not serially correlated.
[23] The test that actual future inflation and expected inflation are cointegrated with zero constant term and unit coefficient is effectively a test of the joint hypothesis of unbiasedness and negligible measurement errors (Engsted, 199 1).
[24] We acknowledge that the selection criterion, the frequency of cointegration across expectations measures, is rather arbitrary.

Table 6 below presents the cointegration vectors based on the normal model. It becomes clear that, for most countries, the slope parameter of the equilibrium relationship deviates significantly from 1, and the intercept significantly from zero. This result corroborates earlier studies which by and large reject the rationality of inflation expectations measures; see, for example, Batchelor and Dua (1987), Evans and Gulamani (1984), Holden and Peel (1977), Pesando (1975), De Menil and Bhalla (1975), De Leeuw and McKelvey (1981), Madsen (1996), Pearce (1979), Pesaran (1985), Thomas (1995) and Figlewski and Wachtel(l98 1).

[insert table 6 here]

Monetary policy analysis

The so-called two-pillar monetary policy strategy of the Eurosystem (where the latter is defined as the ECB and the national central banks of the countries which have adopted the euro as their currency as of January 1, 1999) combines a privileged role for money in the monetary policy decision making process (the first pillar) with a broad-based assessment of prospective inflationary pressures (see Berk, Houben and Kakes, 2000, for details). The latter assessment, or second pillar, implies that the Eurosystem will base its monetary policy decisions also on a host of other indicators for future inflation in the euro area. Measures of inflation expectatations derived from survey data could be useful as information variables under the second pillar. In order to be of value as such there needs to be a stable statistical relationship between expected and future actual inflation. Our finding that current inflation expectations and future realisations of inflation are cointegrated, and that the forecast error is stationary, seems to confirm the usefulness of the former as an information variable for monetary policy: expected inflation derived from consumer surveys shows identical long-run behaviour to the actual inflation 12 months ahead. However, it is not sufficient for our measures of expected inflation to be consistent (or even rational, for that matter) forecasts of actual future inflation. Cointegration, for example, merely implies the existence of a long-term relationship between, in this case, the actual 12-month-ahead inflation rate and expected inflation. It is well-known (Engle and Granger, 1991) that cointegration *per se* does not provide information on the direction of causality in the long-term relationship. The normalization of the cointegrating vector on the inflation rate in table 6 is arbitrary. Yet the direction of causality in the cointegrating relationship between actual future and expected inflation is of considerable importance for monetary policy, for example when the latter is used in feedback rules when setting monetary policy. More generally, if a monetary policy information variable can not be considered (a proxy of) a causal determinant of inflation, the connection between the underlying sources of inflationary pressure (that monetary policy wants to respond to) and the indicator variable is channeled primarily through expectations. In that case it is most plausible that the relation will radically change in the case of a policy intervention that modifies the relation between the underlying states and the inflation that ultimately occurs (Woodford, 1994). In order to further investigate their usefulness for monetary policy purposes, we therefore look for our expectation measures to show a causal relationship with actual future inflation.

A statistical concept which is frequently used to gauge the lines of causality is the traditional Granger causality test, which consist of F-tests on exclusion restrictions in regressions of changes in the (expected) inflation rate. In addition to this test, we investigated the issue of causality by analysing the vector error correction models on which the Johansen cointegration tests reported in table 3 are based. These models are VAR's which include error correction terms consisting of the lagged residual from the cointegrating relations. By Grangers Representation Theorem (Engle and Granger, 1987), the error correction terms provide additional information on the direction of causality.[25] The intuition is that if expected and actual inflation rates have a common stochastic trend, the current change in the latter rate is partly the result of the actual inflation rate moving into alignment with the trend value of the expected inflation rate. Given the normalization in table 6, the expectation measure will Granger cause the actual inflation rate if the error correction term is significant (as measured by conventional *t*-statistics) in the equation for the expected inflation rate, but not in the equation for the actual inflation measure.

[insert table 7]

It follows from table 7 that the traditional Granger causality tests provide, at best, only scant evidence in favour of the hypothesis that the expected inflation rate causes the actual future inflation rate 12 months ahead. Rather, the line of causality seems more often to run from the actual 12-month-ahead-inflation rate to the expected inflation rate, corroborating the findings of Berk (1999). The economic rationale of this seemingly perverse result must be put into

[25] i.e. additional to the traditional Granger causality tests. Note that the latter should be formulated in terms of changes in (expected) inflation because of the persistence.

question, however. This is because the VAR's on which the traditional Granger tests are based use information not available to consumers when forming their expectation of inflation (ie the 12 month-ahead-inflation rate and 11 lags of this rate). The statistics based on the significance of the error correction terms (which are based on information available to consumers when responding to the survey) indicate a somewhat different conclusion. For some large euro area countries which show cointegration between actual future and expected inflation, the hypothesis that causality runs from expected inflation to the actual future inflation rate could not be rejected. The evidence therefore seems to indicate that for Germany, Ireland, Spain, the Netherlands and the euro area, the expectations measures do have predictive power, in a causal sense, for future inflation. This result indicates that our measure of expected inflation for the euro area could enter the second pillar of the monetary policy strategy of the Eurosystem.

Next to being used as an indicator of future consumer price inflation, our expectations measures can, in principle, be used to gauge how consumers' perceptions of future price developments are influenced by certain events relevant to the monetary policy maker. More specifically, we investigate the effects on inflation expectations across European countries of an upturn in past inflation and an unanticipated rise in short-term interest rates. To highlight the monetary policy relevance of this exercise, note that our sample consists of countries such as Germany and the Netherlands, the central banks of which have built up a reputation for credibly holding future inflation to a low stable level, and countries such as for example Greece and Spain, for which (during our sample period at least) no such conclusion could be drawn. Based on the literature on central bank credibility and independence (see, for example, Cukierman, 1994), one might expect, as pointed out by Goodhart (1997), that in the former countries an upturn in inflation has less effect on expectations of future inflation, as the reputation of the central bank prevents the inflationary shock to become persistently embedded in the inflationary expectations of economic agents such as consumers. Similarly, an unanticipated rise in short-term interest rates in countries with more credible central banks should reduce expectations of inflation by more than in countries with less credible central banks. Having available survey-based measures of expected inflation in several countries and over a relatively long time period, it should be possible to put these hypotheses to the empirical test.

Given the preceeding analysis, we explore these hypotheses in a VECM-framework. This way, we take the persistence of both actual and expected inflation into account and at the same time make maximum use of the information provided by the levels of these variables. We first investigated the effects on inflation expectations of European consumers of a change in the most recent actual inflation rate available to them when responding to the survey, that is $\Pi(t-1)$. Table 8 below documents both the long-run reaction of expected inflation (ie the coefficient of the error correction term in the VECM) and the short term reaction (ie the 1-period-effect on expected inflation). As is well known (see Kremers, Ericsson and Dolado, 1992, for details), statistical significance of the former implies cointegration between expected inflation and the one-period lagged inflation rate.[26] The results are disappointing. Most coefficients are insignificant, often of the wrong sign, and do not allow us to discern a pattern between even the polar cases with respect to countries with very credible (Germany) or very incredible (Greece) central banks.

[Insert table 8 here]

A possible explanation of this result is that the changes in inflation had been widely anticipated by consumers. As illustrated by Kuttner (2000), forward-looking expectations should respond only to surprise elements, and not to anticipated movements, in key variables such as the inflation rate. We explore this issue further in a second experiment, in which we analyse the effects of monetary policy surprises on our measures of expected inflation. As a prelude to this experiment, we constructed time series of unanticipated short-term interest rates for the countries in our sample. To this end, we collected data (seasonally unadjusted) on industrial production, a monetary aggregate (Ml because of maximum data availability), and a money market rate (ie 1 month euro rates).*' Based on the same procedure used earlier (see table 1), we could not reject the hypothesis that these data contain a single unit root. We then constructed a five-variable VAR for each country, consisting, in most cases, of the home money market rate, a foreign equivalent, the consumer price inflation rate, industrial production and the money stock. For

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[26] Note that this cointegrating relationship differs conceptually from the ones discussed in tables 3 and 6, which pertained to expected and actual *future* inflation rates.

[27] In contrast to our earlier analysis, we restricted the sample period to run from January 1985 until December 1998, in order to circumvent Lucas (1976)-type of problems due to the shift in monetary policy in most of the countries in our sample as of January 1, 1999. Moreover, we dropped the euro area and the EU from our analysis, because of lack of relevance for our purposes. Portugal was omitted because of data problems.

Germany, the US money market rate was included as the foreign interest rate. For most other European countries, the German short-term interest rate performed this role. Exceptions are the UK, for which both the US and the German rate were included, and Ireland, for which the UK money market rate was included. We then proxied the unexpected short-term interest rate by the residual of the interest rate equation in the VAR. The VAR's proved to be reasonably robust, plausibly signed, and the residuals are white noise processes. The results of this analysis are not shown here in order to save space, but are available from the authors upon request.[28]

Our constructed series of unanticipated movements in the short-term interest rate were included as exogenous variables in a VECM otherwise consisting of expected and actual inflation. We have set up the experiment in such a way that only the most recent actual inflation rate and unanticipated interest rates available to consumers when responding to the survey enter the analysis. The results are presented in table 9.

[Insert table 9 here]

The results are in line with the previous experiment: in all countries under investigation, unanticipated movements in the money market rate fail to elicit statistically significant reactions from our measures of expected inflation.[29] Moreover, the results presented in table 8 are re-confirmed. That is, effects on expected inflation of movements in actual inflation remain murky, at best. Finally, we investigated the relationship between movements in inflation uncertainty reported by consumers (as can be derived from the survey responses) and unanticipated movements in short-term interest rates. It can be argued that the latter elicit relatively small effects on inflation uncertainty in countries with relatively more credible central banks. To test this hypothesis, we ran cross-country regressions using OLS but with an adjustment of the covariance matrix as suggested by Newey and West (1987), as discussed earlier. As can be seen from table 10, we again failed to discover a large number of statistically significant

[28] Note that we have used the VAR's to calculate residuals only, and therefore are not prone to problems of identification or ordering of shocks, as in impulse-response analyses.

[29] We also examined the direct bivariate relationship, estimated with single-equation OLS, between changes in expected inflation and unexpected interest rate movements, with qualitatively similar results.

relationships between movements in inflation uncertainty and unexpected movements in money market rates across countries.[30]

[insert table 10 here]

Tables 8 to 10 therefore seem to suggest that inflation expectations of consumers across European countries, with central banks with varying degrees of credibility, do not in a systematic way, such as for example hypothesised by Goodhart (1997), react to movements in inflation and unanticipated movements in short-term interest rates. These results are however subject to several important caveats. First, they are of course contingent on our constructed measures of expected inflation and unanticipated movements in short-term interest rates. Moreover, the experiments are set up in a very rudimentary way. For example, the modelling of the reaction of consumers' inflation expectations to jumps in inflation and surprise movements in short-term interest rates is rather arbitrary. Investigating the robustness of the results to changes in these assumptions and modelling strategy are important topics for future work.

4 CONCLUDING REMARKS

In this paper we developed and analysed measures of expected future inflation extracted from consumer surveys in the European Union. In these qualitative surveys respondents were asked questions regarding past and future price developments. Six different response categories were allowed for each question. This setup made it possible to develop a variant of the Carlson-Parkin (1977) probability method which generated time-varying response thresholds and did not assume unbiased forecasts. Instead, the expected inflation rate was scaled with the perceived past inflation rate, for which either the actual inflation rate available to respondents when answering the survey, or the information derived from the survey question regarding past price developments was used. We based our probability measures on normal, central and noncentral *t*-distributions. We showed that the actual inflation rate as well as the expected inflation measures were nonstationary, and in most cases were cointegrated. The finding that currently observed inflation expectations of consumers and the unobserved 12-months ahead inflation rate have

[30] This conclusion also holds when we measure inflation uncertainty not by the standard deviation of the inflation rate expected by consumers, but by its coeficient of variation. Results again available on request.

identical long-run properties is of interest for policymakers. But making use of this long-term relationship for monetary policy purposes, for example in the two pillar monetary policy strategy of the Eurosystem, seems hazardous. This is because it remains unclear whether the expectations measures are causal determinants of future inflation. Moreover, and counterintuitively, some experimentation showed that the expectations measures did not react in a systematic way to previous movements in inflation and surprise movements in interest rates. Caution therefore seems to be warranted in using the constructed series of expected inflation for policy purposes.

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Table 1 Unit Root tests on measures of expected inflation
testing the null of 2 unit roots

	normal		symmetrical-t		asymmetrical-t		actual
	actual	perceived	actual	perceived	skew 1	skew 2	inflation
Belgium	6.13**	7.76”	6.17”	8.35”	5.68**	8.07”	5.98”
Germany	6.60”	8.23**	6.56**	8.27**	4.72”	7.54”	5.51”
France	6.11**	5.83”	6.09**	5.82**	5.32”	6.29”	6.24**
Ireland	6.48**	9.43**	6.60”	9.09**	4.26**	8.02**	5.00**
Italy	6.04**	6.76**	6.02**	7.22**	4.16**	6.61**	3.78**
Netherlands	6.28**	6.75**	6.23**	7.04**	6.61**	8.62**	5.58**
Spain	5.51**	7.06”	5.55”	7.25**	6.72**	7.17**	7.05**
Portugal	4.71**	7.81*+	4.73**	8.16**	5.50**	7.26**	5.17**
Euro area	5.49**	5.85**	5.49**	5.82”	4.28**	6.47**	4.98**
Denmark	6.40”	6.64**	6.39**	6.96”	6.67**	9.63**	6.74**
Greece	4.57**	7.32**	4.55**	7.45**	4.36**	7.23**	4.97**
United Kingdom	4.44**	7.86**	4.39**	8.60**	3.99**	8.02”	3.47**
EU	5.70	6.10**	5	6	6.40**	8.93**	5.6

Notes: expectations measures are based on lagged actual inflation rate or the perceived inflation rate derived from survey. The asymmetric distribution uses skewness based on either information available when responding to the survey (skew 1) or the entire sample (skew 2). See text for details.
Figures presented indicate absolute values of the ADF-test, based (unless indicated otherwise) on a deterministic process without intercept and linear trend, and with 3 lagged differences.(x,y) specifies the deterministic process, with x=trend and intercept t, intercept c or none n, and y the number of lagged differences. ** (*) indicates significance at 1% (5%).

Table 1 Unit Root tests on measures of expected inflation
testing the null of 1 unit foot

	normal		symmetrical-t		asymmetrical-t		actual
	actual	perceived	actual	perceived	skew 1	skew 2	inflation
Belgium	2.82	3.41’	2.82	2.6	2.72	2.08	3.00 (t,3)
Germany	0.96	1.31	0.97	1.38	2.37	1.47	1.49
France	2.14	2.63	2.18	2.62	4.10**	2.87	1.3
Ireland	1.86	1.86	1.85	1.93	2.51	2.39 (t,12)	2.02
Italy	0.69	0.68	0.67	0.66	1.83 (c,12)	1.14	1.54
Netherlands	1.58	1.39	1.58	1.46	1.77	1.93	1.69
Spain	1.98	1.86	1.96	1.73	2.86	2.07	2.23
Portugal	1.3	1.96 (c,12)	1.27	2.13 (c,12)	1.99 (c,12)	1.45	0.55
Euro area	0.95	1.31	0.95	1.28	2.08	2.7	0.95
Denmark	3.12’	3.41*	3.20*	2.79	3.18	4.49** (t,3)	2.63
Greece	1.95	1.91	1.97	2.01	2.51 (c,12)	1.72	1.93
United Kingdom	1.6	1.36	1.61	1.62	2.15 (c,12)	2.58	1.45
	1.22						
EU	1.22	1.66	1.22	1.54	1.91	2.67	1.52

Notes: expectations measures are based on lagged actual inflation rate or the perceived inflation rate derived from survey. The asymmetric distribution uses skewness based on either information available when responding to the survey (skew 1) or the entire sample (skew 2). See text for details.
Figures presented indicate absolute values of the ADF-test, based (unless indicated otherwise) on a deterministic process with intercept, and including 3 lagged differences.(x,y) specifies the deterministic process, with x=trend and intercept t, intercept c or none n, and y the number of lagged differences.
** (*) indicates significance at 1% (5%).

Table 2 Unit Root tests on measures of expected inflation uncertainty
testing the null of 2 unit roots

	normal		svmmetrical-t		asvmmetrical-t	
	actual	perceived	actual	perceived	skew 1	skew 2
Belgium	6.12**	8.04**	6.16**	8.49"	6.22"	6.21•*
Germany	6.10**	7.90**	6.11**	8.33**	6.45**	6.11"
France	6.34"	6.18**	6.20"'	5.93**	6.33**	6.31•*
Ireland	5.20"	7.02"	5.28**	6.70**	5.44"	5.29**
Italy	5.27**	6.64"	5.17**	6.68**	5.65**	5.32**
Netherlands	6.13**	6.84**	6.18**	6.73**	6.09**	6.18**
Spain	6.46**	7.87"	6.56**	8.39"	6.79**	6.61**
Portugal	5.43"	8.22**	5.75"	8.61**	5.80**	5.75**
Euro area	5.54**	6.37"	5.52**	6.46**	5.69**	5.63**
Denmark	6.94**	7.63**	6.97**	8.10**	7.09**	7.33**
Greece	5.10"	6.53"	5.29**	6.24**	6.13**	4.96**
United Kingdom	4.06**	8.76**	4.15**	9.13"	4.09**	4.06**
EU	6.90**	6.66"	[illegible]	06"	5"	0 "

Notes: expectations measures are based on lagged actual inflation rate or the perceived inflation rate derived from survey. The asymmetric distribution uses skewness based on either information available when responding to the survey (skew 1) or the entire sample (skew 2). See text for details.
Figures presented indicate absolute values of the ADF-test, based (unless indicated otherwise) on a deterministic process without intercept and linear tend and 3 lagged differences.(x,y) specifies the deterministic process, with x=trend and intercept t, intercept c or none n, and y the number of lagged differences.
** (*) indicates significance at 1% (5%).

Table 2 Unit Root tests on measures of expected inflation uncertainty
testing the null of 1 unit root

	normal		svmmetrical-t		asvmmetrical-t	
	actual	perceived	actual	perceived	skew 1	skew 2
Belgium	2.82	3.23 (t,3)	2.77	2.81 (c,4)	2.69	2.81
Germany	1.28	1.6	1.33	1.96	1.6	1.31
France	2.48	2.84	2.55	3.03 (t,3)	2.53	2.56
Ireland	2.37	2.66	2.34	2.81	2.56	2.32
Italy	2.46	2.78	2.47	2.43	2.52	2.47
Netherlands	1.51	2.34	1.51	2.49	1.54	1.51
Spain	1.97	1.79	1.95	1.58	2.09	1.88
Portugal	1.03	2.07 (c,12)	1.04	2.57 (c, 12)	1.6	0.99
Euro area	1.05	2.14	1.08	2.64	1.22	1.07
Denmark	2.72	2.15	2.69	1.65	2.74	2.69
Greece	2.28	1.99	2.34	1.89	1.81 (c,12)	2.37
United Kingdom	1.7	1.6	1.68	2.47	2.06	1.72
EU	0.8	1.91	0.76	3 17	0.81	0.75

Notes: expectations measures are based on lagged actual inflation rate or the perceived inflation rate derived from survey. The asymmetric distribution uses skewness based on either information available when responding to the survey (skew 1) or the entire sample (skew 2). See text for details.
Figures presented indicate absolute values of the ADF-test, based (unless indicated otherwise) on a deterministic process with intercept, and including 3 lagged differences.(x,y) specifies the deterministic process, with x=trend and intercept t, intercept c or none n, and y the number of lagged differences.
• * (*) indicates significance at 1% (5%).

Table 3 Testing for cointegration between expected and actual inflation

dimension of VAR in parentheses

	normal		svmmetrical-t		asvmmetrical-t	
	actual	perceived	actual	perceived	skew 1	skew 2
Belgium	Y (12)	Y (12)	Y (12)	Y (12)	Y (12)	Y (12)
Germany	Y (12)	Y (3)	Y (12)	Y (3)	Y (12)	N (3)
France	N (3)	N (3)	N (3)	N (12)	N (3)	N (3)
Ireland	Y (12)	N (6)	Y (12)	N (6)	N (12)	N (6)
Italy	N (3)	N (3)	N (3)	N (3)	N (3)	N (3)
Netherlands	Y (6)	Y (9)	Y (6)	Y (9)	N (6)	Y (6)
Spain	Y (12)	N (3)	Y (12)	N (12)	N (9)	Y (12)
Portugal	N (3)	N (3)	N (3)	N (3)	N (3)	N (3)
Euro area	Y (12)	N (3)	Y (12)	N (9)	N (12)	N (3)
Denmark	X	X	X	X	X	X
Greece	Y (3)	N (3)	Y (3)	N (3)	N (6)	N (3)
United Kingdom	Y (12)	Y (9)	Y (12)	Y (9)	N (9)	N (9)
FU	Y (12)	Y (3)	Y (12)	Y (3)	N (6)	N (12)

Notes: expectations measures are based on lagged actual inflation rate or the perceived inflation rate derived from survey. The asymmetric distribution uses skewness based on either information available when responding to the survey (skew 1) or the entire sample (skew 2). See text for details.
Y (N) indicate presence (lack) of cointegration at the 5% significance level.Cointegration tests assumed no trend in data, but included and intercept in cointegration equation.The dimension of the VAR is based on AIC, SBC and LR tests. X indicates not applicable due to different orders of integration.

Table 4 Testing for stationarity of the forecast error

testing the null of 2 unit roots

	normal		svmmetrical-t		asvmmetrical-t	
	actual	perceived	actual	perceived	skew 1	skew 2
Belgium	6.74**	7.80**	6.77**	7.91**	5.62**	7.84"
Germany	6.30**	7.23**	6.29**	7.20**	5.17**	X
France	X	X	X	X	X	X
Ireland	4.76**	X	4.80"	X	X	X
Italy	X	X	X	X	X	X
Netherlands	7.03**	6.90"	6.97"	7.26"	X	8.88**
Spain	5.95"	X	5.96**	X	X	6.30**
Portugal	X	X	X	X	X	X
Euro area	5.52**	X	5.52**	X	X	X
Denmark	X	X	X	X	X	X
Greece	4.93**	X	4.93**	X	X	X
United Kingdom	4.24**	5.80**	4.21**	5.93"	X	X
FU	6.80**	0**	6.8 **	2**	X	X

Notes: expectations measures are based on lagged actual inflation rate or the perceived inflation rate derived from survey. The asymmetric distribution uses skewness **based on** either information available when responding to the survey (skew 1) or the entire sample (skew 2). See text for details.
Figures presented indicate absolute values of the ADF-test, with a deterministic process without intercept and/or trend and with 3 lagged differences. and y the number of lagged differences, • * (*) indicates significance at 1% (5%). X indicates not applicable because of lack of cointegration.

Table 4 Testing for stationarity of the forecast error
testing the null of 1 unit root

	normal		symmetricai-t		asymmetrical-t	
	actual	perceived	actual	perceived	skew 1	skew 2
Belgium	4.10** (c,12)	4.07** (c, 12)	4.08** (c, 12)	3.83** (c, 12)	4.72** (c, 12)	3.31• (c,12)
Germany	2.92** (n, 1)	2.34' (n,1)	2.91** (n,1)	2.21* (n,1)	2.19* (n,1)	X
France	X	X	X	X	X	X
Ireland	4.04" (c, 12)	X	4.02** (c,12)	X	X	X
Italy	X	X	X	X	X	X
Net herlands	2.91* (n,12)	3.33' (c,8)	2.89" (n,12)	3.33' (c,8)	X	5.03** (c, 12)
Spain	3.09' (c,1)	X	3.09' (c, 1)	X	X	3.35 (t,0)
Portugal	X	X	X	X	X	X
Euro area	2.03* (n,1)	X	1.99' (n,1)	X	X	X
Denmark	X	X	X	X	X	X
Greece	2.03* (n,1)	X	2.02* (n,1)	X	X	X
United Kingdom	2.31* (n,1)	1.92 (n,1)	2.32' (n, 1)	1.77 (n,1)	X	X
FU	[illegible]	9 [illegible]	99'	7 75 (c,1)	X	X

Notes: expectations measures are based on lagged actual inflation rate or the perceived inflation rate derived from survey. The asymmetric distribution uses skewness based on either information available when responding to the survey (skew 1) or the entire sample (skew 2). See text for details. Figures presented indicate absolute values of the ADF-test.(x,y) specifies the deterministic process, with x=trend and intercept t, intercept c or none n, and y the number of lagged differences. ** (*) indicates significance at 1% (5%). X indicates not applicable because of lack of cointegration.

Table 5 Summary statistics of forecast error for selected expectations measures

	MAPF		RMSE		Theil	
	NO	TO	No	TO	NO	TO
Belgium	1.02	1.02	1.32	1.32	0.21 (5.59)	0.21 (5.60)
Germany	1.06	1.06	1.27	1.27	0.11 (4.52)	0.11 (4.52)
France	X	X	X	X	X	X
Ireland	1.01	1.03	1.31	1.33	0.14 (4.37)	0.14 (4.43)
Italy	X	X	X	X	X	X
Netherlands	0.89	0.89	1.07	1.07	0.18 (6.02)	0.18 (6.02)
Spain	1.66	1.68	2.02	2.05	0.06 (4.37)	0.06 (4.43)
Portugal	X	X	X	X	X	X
Euro area	0.88	0.89	1.1	1.11	0.09 (15.33)	0.09 (15.52)
Denmark	X	X	X	X	X	X
Greece	4.17	4.14	4.98	4.95	0.02 (2.69)	0.02 (2.67)
United Kingdom	1.68	1.68	2.33	2.31	0.07 (6.02)	0.07 (5.99)
EU	0.89	0.91	1.11	1.12	0.08 (17.75)	0.08 (17.92)

Notes: NO (TO): expectations measures based on normal (symmetric-t) distributions and the lagged actual inftation rate.
MAPE: mean absolute prediction error, RMSE: root mean square error, Theil: Theil's inequality coefficient vis-a-vis assumption of unchanged price level (unchanged inflation rates). X indicates not applicable because of lack of cointegration.

Table 6 Cointegrating vectors for inflation expectations based on **normal** model

	expected inflation	actual inflation 13 month ahead	intercept
Belgium	1	-1.37 (0.08)	1.44 (0.19)
Germany	1	-1.21 (0.04)	0.74 (0.12)
France	X	X	X
Ireland	1	-0.80 (0.03)	0.46 (0.10)
Italy	X	X	X
Netherlands	1	-0.83 (0.20)	-0.18 (0.43)
Spain	1	-0.87 (0.04)	1.31 (0.23)
Portugal	X	X	X
Euro area	X	X	X
Denmark	X	X	X
Greece	1	-1.03 (0.17)	0.22 (2.43)
United Kingdom	1	-1.22 (0.03)	1.19 (0.13)
EU	X	X	X

Notes: presented are cointegrating vectors on inflation of the form [expected actual intercept] with standard errors in parentheses. The dimension of the VAR on which these relations are based is as in table 3. X= NA

Table 7 Causality tests

	$\pi^e \neq \pi$		$\pi \neq \pi^e$	
	F test	t-test on FCM	F test	t-test on ECM
Belgium	1.33	7.88""	4.53""	2.31**
Germany	2.12*	7.18""	2.68""	0.76
France	1 .64**	X	2.58""	X
Ireland	0.80	9.64""	1.89**	1.69
Italy	1.33	X	2.39""	x
Netherlands	1.48	4.47""	1.11	0.25
Spain	0.87	6.08""	2.63**	0.35
Portugal	2.15"	X	1.24	X
Euro area	3.81**	3.87""	5.80""	1.45
Denmark"	4.60""	X	7.78""	X
Greece	2.41**	4.95""	2.14"	3.16""
United Kingdom	3.45**	8.81**	5.10**	0.05""
EU	2.28**	5.62**	5.75**	2.23**

Note VECM and traditional Granger causality test include 12 lags.
° VAR in levels, 12 lags. Investigated is causal relationship between actual future inflation rate (12 months ahead) and expected inflation rate. **(*) indicates significance at 1 % (5%)

Table 8 Effect on expected inflation of change in actual inflation

	lona-run (ECM)	short term (1 period)
Belgium	0.16"	0.15
	(1.99)	(1.33)
Germany	-0.03	-0.16
	(0.3)	(0.90)
France	0.24**	-0.32"
	(3.24)	(2.20)
Ireland	-0.29 **	-0.05
	(3.37)	(0.43)
Italy	0.04	0.26
	(1.09)	(0.98)
Netherlands	-0.13 *	0.21
	(2.4)	(1.66)
Spain	-0.17 *	0.04
	(1.98)	(0.32)
Portugal	-0.15"	0.01
	(2.55)	(0.03)
Euro area	-0.06	-0.10
	(1.10)	(0.67)
Denmark	X	X
Greece	-0.01	-0.07
	(0.20)	(0.46)
United Kingdom	0.17	0.38
	(1.46)	(1.77)
EU	-0.12	-0.07
	(1.70)	(0.42)

Notes: Presented are effects of change in lagged actual inflation rate, estimated with VECM or VAR, including 12 lags.
absolute t-values in parentheses. **(*): significant at 1% (5%).
X=NA. Sample: 1986: l-l 999: 12

Table 9 Effects on expected inflation rate of change in:

	actual inflation rate		unexpected interest rate
	lona-run (ECM)	short term (1period)	
Belgium	0.08	-0.03	-0.01
	(1.34)	(-0.46)	(0.06)
Germany	-0.10	-0.10	-0.13
	(1.53)	(-0.95)	(0.58)
France	-0.14"	0.10	0.00
	(2.02)	(0.68)	(0.03)
Ireland	-0.05	-0.05	-0.01
	(0.47)	(-0.58)	(0.53)
Italy	-0.04	-0.1 1	-0.01
	(0.95)	(0.50)	(0.08)
Netherlands	-0.08	-0.25	0.01
	(1.42)	(I .87)	(0.02)
Spain	-0.14	0.36**	-0.06
	(1.56)	(3.10)	(0.51)
Portugal	X	X	X
Euro area	X	X	X
Denmark	X	X	X
Greece	-0.16"	0.14	-0.02
	(-2.96)	(I .36)	(0.37)
United Kingdom	-0.14	0.15	0.14
	(-1.44)	(0.91)	(0.26)
EU	X	X	X

Notes: Presented are effects of change in lagged actual inflation rate and in unexpected short-term interest rate, estimated with an VECM/VAR including 12 lags, and treating the unexpected interest rate exogenous. See main text for discussion on the construction of unexpected interest rate. Absolute t-values in parentheses. **(*): significant at 1 % (5%). X=Na.
Sample: 1987:1-1998:12

Table 10 Effects on inflation uncertainty of unexpected movement in short-term Interest rates

	intercept	unexp int rate
Belgium	-0.00	-0.04
	(0.15)	(0.56)
Germany	0.01	0.03
	(0.40)	(0.20)
France	-0.01	0.02
	(0.58)	(0.30)
Ireland	-0.00	-0.00
	(0.22)	(0.23)
Italy	-0.01	0.07
	(0.24)	(0.91)
Netherlands	0.01	0.20
	(0.39)	(0.70)
Spain	-0.02	0.01
	(1.29)	(0.06)
Portugal	X	X
Euro area	X	X
Denmark	X	X
Greece	-0.04	0.01
	(0.67)	(0.26)
United Kingdom	-0.00	-0.08
	(0.08)	(0.71)
EU	X	X

Notes: movements in inflation uncertainty measured by change in standard deviation as derived from survey. Absolute t-values, computed with Newey-West standard errors, in parentheses.
X= not available